# Mandated Access and the Make-or-Buy Decision: The Case of Local Telecommunications Competition

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an empirical question cannot be settled by non-empirical arguments.
George Stigler, *The Organization of Industry* (1968), p. 115.

Abstract: When the facilities of an incumbent monopolist are made available to potential competitors through some type of "essential facilities" or related claim, a comnion concern is that the ability to "buy" inputs substantially attenuates the incentive to "make" inputs. In this paper, we evaluated both theoretically and empirically the relationship between "make" and "buy." In our particular construct, three sometimes-conflicting effects are relevant to the "make-or-buy" decision, of which the substitution effect is only one. Our empirical example considers the deployment of switching facilities by entrants to local exchange telecommunications markets, and these empirics indicate that the substitution effect is not dominant. While particular to telecommunications, our findings do support the general notion that the substitution effect is not the only relevant consideration, either theoretical or empirical, for policy makers in selecting what inputs to make available to entrants when promoting competition in the utility industries.

## I. Introduction

Over the past decade or so, considerable attention has been directed to the promotion of competition in and the eventual deregulation of the public utilities -- gas, electricity, and local telecommunications. As part of this effort, potential competitors often are given access to elements of the incumbent monopolist's network or plant.[1] Such access is required when particular elements of the

[1] In some cases. such as local telecommunications, tlie incumbent continues to provide retail services so that the entrants are both competitors and customers (or "competitor customers") of thr incunihent. 111 others, such as electricity. the incumbent often is prohibited from

incumbent network continue to possess natural monopoly characteristics such as sizeable scale and scope economies.[2] Whether access to these elements is based on the theory of "essential facilities" of antitrust or "unbundled network elements ("UNEs")" of the Telecomniuiiications Act of 1996, the result is the same: entrants are allowed to use the facilities **of** the incumbent as their own, and such access **is** priced at some measure of "cost," typically some variant of forward-looking economic cost.

A principle difficulty faced by policy makers in this context is which elements of the network are "essential facilities" or satisfy some other governing standard such as the "impairment" standard of Section 251(d)(2) of the 1996 Telecommunications Act.[3] Economists and lawyers have described numerous problems with both the over- and under-inclusion of elements within the (broad) category of "essential." One frequent concern, particularly in the debate over local exchange telecommunications competition, is that by giving entrants access to parts of the network, those components of the network will never be duplicated and thus subject to the competitive pressure required to deregulate. Areeda and Hovencamp (1996, ¶ 771) observe, "the right to share a monopoly discourages firms from developing their own alternative inputs, ... [though, t]o be sure, that incentive may not be removed altogether." This substitution effect, commonly couched in terms of a "make-or-buy" decision by the entrant, often lies at the core of the arguments by those calling for a less inclusive policy on what is or is not "unbundled" in modern telecommunications competition policy.

With respect to post-1996 Act telecommunications policy, the courts have been somewhat schizophrenic on the question of unbundling network elements and the incentive of competitors to vertically integrate by supplying their own inputs. In affirming the FCC's cost standard for pricing unbundling elements, the

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participating in the inarket targeted for competition and deregulatioii (whether upstream or downstream). Mandated access to AT&T's network was an important driver in the development of competition in the U.S. long **distance** industry. See Brock (1981)and Cantelon **(1993).**

[2] Such supply-side cliaracteristics are prevalent in the more geographically local elements of the aforementioned utilities' plant.

[3] Section 251(d)(2)(B) **of** the 1996 Telecommunications Act requires the FCC in determining what network elements should be made available to consider, at a minimum, whether "the failure to provide access to such **network** elements would impair the ability of the telecommunications **carrier** seeking access to provide **the** services that it **seeks** to offer." **47** U.S.C. § 251(d)(2)(B). The Telecommunications **Act** also contains a "necessary standard" in § 251(d)(2)(A) -- that is, providing access to any "proprietary" network element must be necessary for **the** requesting carrier to provide service. In practice, the necessary standard is rarely relevant

Supreme Court, in *Verizon v. FCC*, 122 S.Ct. 1646 (2002) responded to the plaintiff's claim that unbundling did not stimulate facilities-based entry by observing, "actual investment in competing facilities since the effective date of the Act simply belies the no-stimulation argument's conclusion (at 1669)."[4] Less than a fortnight after the Court issued its Opinion in *Verizon*, the D.C. Circuit in *USTA v. FCC*, 290 F.3d 415 (D.C. Cir. 2002) responded, "the existence of investment of a specified level tells us little or nothing about incentive effects. The question is how such investment compares with what would have occurred in the absence of the prospect of unbundling [citation omitted], an issue on which the record appears silent. Although we can't expect the [Federal Communications] Commission to offer a precise assessment of disincentive effects (a lack of multiple regression analyses is not ipso facto arbitrary and capricious), we can expect *at* least some confrontation of the issue and some effort to make reasonable trade-offs." (at 425). The Appeals Court also recognized, "access to UNEs may enable a CLEC to enter the market gradually, building a customer base up to the level where its own investment would be profitable." (at 424). This latter observation echoes Areeda and Hovencamp, who conclude, "the plaintiff may begin building a customer base and might eventually acquire enough customers to build its own pipeline but would not have done so if not permitted to enter the market by sharing (1996, ¶ 771)." Obviously, it is difficult to establish an unambiguous relationship between access of rivals to the monopolist's network and the incentive of entrants to construct their own inputs. In the end, the question is empirical.

The purpose of this paper is to evaluate in both a theoretically and empirically rigorous way the issue of how regulated access to inputs influences a firm's incentives to vertically integrate in order to self supply such inputs. Theoretically, the presence of a substitution effect is undeniable. However, the theory reveals **two** other effects, one working with (the *scale effect*) and the other against (the *entry effect*) the substitution effect.[5] Which of the three effects dominates cannot be determined solely by theory. Consequently, an empirical test of the theory is conducted, with the deployment of switching equipment by

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[4] The Telecommunications Act requires that network access, or unbundled elements ("UNEs") be price at "cost." Cost was to be defined by the Federal Communications Commission, and that agency adopted a total-element, long-run incremental cost ("TELRIC") cost standard. TELRIC is a forward-looking methodology, where costs are based on the most efficient, currently deployed technology. See *In re Implementation of the Local Competition Provisions of the Telecommunications Act of* 1996, First Report and Order, FCC No. 96-325, 11 FCC Rcd 15499 (rel. Aug. 8, 1996).

[5] There are, no doubt, many other ways in which this issue can be evaluated in a theoretical context, which may point to additional effects that produce ambiguity.

competitive local exchange carriers ("CLECs")as a case study.6 This case study is particularly relevant to this issue, given that the entrant's access or lack thereof to the switching function of the local exchange network is the subject of heated debate (Sunderland; Bischoff). The empirical results indicate that for this particular case, the substitution effect is not dominant; restricted access to the "switching element" of the local exchange access, either through higher prices are outright restrictions, will discourage switching facilities deployment by entrants.

The empirical findings of this paper provide important guidance **for** competition policy in the local exchange telecommunications market. Indeed, at the heart of the current telecommunications policy debate lays a key unanswered question: what public policy will best promote facilities-based entry into the local exchange telecommunications marketplace?[7] **At** the center of the debate is the question as to whether the requirement of the 1996 Telecommunications Act that incumbent local telephone carriers ("ILECs") provide access to their local networks to new entrants, or the requirement that such access be made available at "cost," promotes or deters facilities-based entry. The ILECs encourage policy makers to limit access to their network (particularly unbundled switching) and, when access is provided, that it be priced high. Without access to the incumbent's network or with access only at high prices, the ILECs contend that CLECs will be forced to deploy their own facilities and consequently will do so. In other words, the ILECs implicitly assume there exist a strong substitution effect between access to the existing network and the construction of new network. 'The CLECs, the Federal Communications Commission ("FCC"), and Congress generally disagree with the ILECs claims, requiring the ILECs to unbundled their networks and make these components available to retail rivals. While the debate over unbundled elements does not lack of verve, what is missing from the debate is any semblance of a theoretical framework within

[6] This analysis responds directly to the D.C. Appeals Court's desire for "multiple regression analyses."

[7] It is unclear why the debate focuses on this question from **a** policy perspective, since the 1996 Telecom Act indicates no preference for facilities-based competition over competition using unbundled element. Indeed, Section 271 requires the existence ***of*** both forms of toinpetition *in* order for a Bell Company to offer interLATA long distance services. From the perspective of the incumbent monopolist, however, the entry-deterring effect of forcing rivals to incur the sunk costs of network deployment is plainly desirable.

which to analyze the issues and, perhaps more disturbing, any empirical evidence.[8] We attempt to address these two shortcomings in this paper.

This paper is organized as follows. In Section II, a two-stage, game-theoretic model of switch deployment is presented. This theoretical analysis, though simple, illustrates the difficulty in finding an unambiguous relationship between network access prices and CLEC facilities deployment. In Section III, the empirical model is described and the results summarized. Concluding comments are provided in Section IV.

## 11. Conceptual Framework

In order to assess the impact of wholesale prices for loops and switching on switch deployment, we develop an economic model in the form of **a** two-stage game. In Stage 1, firms choose whether or not to enter the market. Then, in Stage 2, firms choose how much switching to self-supply. **As** is customary with two-stage models, the model is solved backwards so that the first decision to evaluate is how a firm selects its optimal investment in switching, $S$, given that it enters in Stage 1. For simplicity, it is assumed that firms are symmetric *ex ante,* hut not *ex post,* and that entry does not affect the retail margin.

'The model takes the point of view of the CLEC and evaluates the CLEC's decision whether or not to self-provide local switching. In other words, the model assumes that this CLEC entrant decides on its switch investment prior to knowing how many customers it will have (i.e., prior to entry). Thus, there is an uncertainty component to the model, and this uncertainty relates to demand. Upon entering the market, the CLEC provides service to end-users using unbundled loops purchased from the ILEC along with either unbundled local switching purchased from the ILEC or its own, self-supplied local switching.

The variables of the model include:

| | |
|---|---|
| $I$ = | the number of firms that enter; |
| $N(I)$ = | expected number **of** customers a single firm acquires and serves upon entry; |

[8] Two empirical studies address the impact of the FCC's restriction on unbundled switching in the largest metropolitan statistical areas. See Z-Tel Communications (2002a, 2002b). Neither of these papers addresses, however, the question of facilities-deployment and network access prices.

$\lambda N(I)$ = actual number of customers;

$\lambda$ = random variable, $E(\lambda) = 1, \lambda \in [0, \infty +)$ with probability density function $f(\lambda)$ and cumulative density function $F(\lambda)$;

$S$ = number of customers firm can service with its own switches;

$e \cdot S$ = cost of firm switches (a sunk cost), where $e$ is the price per customer served by self-supplied switching;

$P_l$ = regulated price of an unbundled loop;

$P_s$ = regulated price of unbundled switching;

$c$ = other per customer retail costs;

$R$ = revenue per end-user customer;

$M_0$ = margin with self-supplied switching $(R - P_l - c)$;

$M_b$ = margin with unbundled switching $(R - P_l - P_s - c)$, where $M_0 > M_b$.

Prior to entry, firms expect to acquire and serve N customers. However, the customer base is only an expectation, with actual customers equaling $\lambda N$ (where $\lambda$ is a random variable). If $\lambda N < S$, actual demand is less than switching capacity, the entrant uses its own switching exclusively. This level of demand occurs with probability $F(S/N)$.

In this case, the profit of the entrant is

$$\pi = \lambda N \cdot M_0 - e \cdot S, \tag{1}$$

which is simply the margin on the actual customer base minus switch investment. Alternately, if $\lambda N > S$, the entrant **uses** both its own switching capacity as well as purchasing unbundled switching from the ILEC. This level of demand occurs with probability $[1 - F(S/N)]$. In this case, the **profit** of the entrant is

$$\pi = S \cdot M_0 + (\lambda N - S) M_b - e \cdot S. \tag{2}$$

Note that there can be other sunk entry costs in addition to switching investment, but the presence of such costs does not alter the analysis. For expositional convenience, we ignore such costs.

Expected profit as a function of S, N, $P_l$, and $P_s$ is

$$E\pi = \int_0^{S/N} \lambda f(\lambda) d\lambda \;\; N \cdot M_o + \int_{S/N}^{\infty} \lambda f(\lambda) d\lambda \cdot NM_b + (1 - F(S/N)) \;\; S \cdot (M_o - M_b) - e \cdot S \,. \tag{3}$$

To find the optimal level of switch investment, $S^*$, the first order condition of Equation (3) with respect to $S$ is needed:

$$\frac{\partial E\pi}{\partial S} = (1 - F(S/N)) \cdot (M_o - M_b) - e = 0 \,. \tag{4}$$

The second order condition is

$$\frac{\partial E\pi}{\partial S} = -f(S/N) \cdot (1/N) \cdot (M_o - M_b) < 0 \tag{5}$$

indicating that $S^*$ **is** a maximum.

### 1. COMPARATIVE STATICS

Useful comparative static results include

$$\frac{\partial^2 E\pi}{\partial S \partial N} = -f(S/N) \cdot \frac{-S}{N^2}(M_o - M_b) > 0, \tag{6}$$

indicating that the larger the number of expected customers, the more the entrant will self-supply switching. Defining $\pi$ at $S^*$ as $\pi^*$, we have

$$\frac{\partial E\pi^*}{\partial N} = \int_0^{S/N} \lambda f(\lambda) d\lambda \cdot N \cdot M_o + \int_{S/N}^{\infty} \lambda f(\lambda) d\lambda \cdot NM_b > 0 \,, \tag{7}$$

$$\frac{\partial E\pi^*}{\partial P_s} = N \left[ (1 - F(S/N)) \cdot S/N - \int_{S/N}^{\infty} \lambda f(\lambda) d\lambda \right] < 0 \,, \tag{8}$$

and,

(9)

Equation **(7)** indicates that an increase in the customer base increases expected profits. Equation (8) and Equation (9) imply that higher wholesale prices for loops or switching reduce expected profits.

Turning to the question of switches deployed in the market, assume that all firms pick the same $S^*$ *ex ante*, but ***ex** post* the demands differ randomly across firms. Market demand is assumed to be constant and insensitive to the allocation of demand among firms. Given $R$, $P_l$, $P_s$, e, and $N$, each firm selects $S^*$. Equilibrium profit for each firm, $\pi^*$, **is** assumed to be zero. This assumption allows us to solve for $N$, the "minimum necessary market size" or "minimum viable scale ("MVS")."[9] The number of firms that enter, I, depends on this N (*i.e.*, $I = I(N)$), where $I' < 0$ -- the larger the market share needed to break even (i.e., the larger is MVS), the fewer firms enter in equilibrium.[10] The optimal level of switch deployment for any given firm is $S^* = S^*(P_l, P_s, N)$.

If each firm deploys $S^*$ switching, then the total amount of CLEC switching is given by

$$\breve{S} = I(\breve{N}) \cdot S^*, \tag{10}$$

which states that total switching capacity deployed is simply the number of firms multiplied by average switching capacity.

*Switch Deployment and Loop Prices*

The response of switching deployed to a change in the loop rate is

$$\frac{d\breve{S}}{dP_l} = I' \cdot \frac{\partial \breve{N}}{\partial P_l} \cdot S^* + I\left[\frac{\partial S^*}{\partial P_l} + \frac{\partial S^*}{\partial N}\frac{\partial \breve{N}}{\partial P_l}\right] \tag{11}$$

but $dS^*/dP_l = 0$, so

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[9] Minimum **viable** scale **is** the "smallest annual level of **sales** that the committed entrant must **persistently** achieve **for profitability**" *Horizontal* **Merger** *Guidelines* § 3.3.

[10] For an excellent analysis **of entry dnd sunk** *costs* **generally, see Sutton** (1991)

$$\frac{d\check{S}}{dP_l} = \frac{\partial \check{N}}{\partial P_l}\left[ I' \cdot S^* + I \cdot \frac{\partial S^*}{\partial N} \right] \tag{12}$$

All the right-hand side terms in Equation (12) are positive except for $I'$. Thus, the sign on $d\check{S}/dP_l$ is ambiguous. Equation (12) reveals the two important, and contrary, effects of changes in the loop rate on switch deployment. First, as $P_l$ rises, the per-customer margin declines. When customers become less profitable, the entrant needs more customers to breakeven ($dN/dP_l > 0$), and an increase in customers leads to increased switch deployment. This effect is called the *scale effect,* which arises simply from the fact that the smaller are per-unit profit margins, the larger is minimum viable scale.

The second effect is called the *entry effect* (the bracketed term in Eq. 12). From the scale effect, we know that a change in the loop price alters the scale of the firm. **As** the market share required to profitably enter rises due an increase in the loop rate, fewer firms can profitably enter ($I' < 0$). **A** reduction in the number of firms reduces total switch deployment, ceteris paribus. The ambiguous relationship between input price and switch deployment arises from the fact that the entry effect opposes the scale effect and it is impossible to determine from theory which effect dominates.

Intuitively, the source of ambiguity can be described simply as follows. The total quantity of switching deployed ($\check{S}$) equals the amount of switching deployed by each firm ($S^*$) multiplied by the number of firms (N). Higher loop prices raise MVS, thereby reducing the number of firms but increasing the quantity of switching deployed by each of the remaining firms. No unambiguous claim can be made about the product of the two quantities ($S^*$ and $N$) at different loop prices.

*Switch Deployment and Switching Prices*

While the scale and entry effects arise when considering the effects of the switching price on total switches, an additional effect is also present. **A** change in the switching rate on total switches is

$$\frac{d\check{S}}{dP_s} = \frac{\partial \check{N}}{\partial P_s}\left[ I' \cdot S^* + I \cdot \frac{\partial S^*}{\partial N} \right] + I \cdot \frac{dS^*}{dP_s} \tag{13}$$

The scale and entry effects are both present, but there is an additional term on the right-hand side not present in Equation (12). This term measures the *substitution effect.* The substitution effect accounts for the substitution between self-supplied switching and purchased switching. **As** the price of purchased switching declines, the incentive to self-supply switching declines ($dS^*/dP_s > 0$), and vice versa. Clearly, the substitution effect is only one of three potential effects arising from a change in switching rates. The sign of Equation (13), as with Equation (12), is ambiguous.

### 2. FROM THEORY TO EMPIRICS

Equations (12) dnd (13), while failing to establish a definite sign for the relationship between the relevant prices and switch deployment, highlight the factors involved and allow us to formulate an appropriate econometric model. To summarize, both loop rates and unbundled switching rates have ambiguous effects on the deployment of switches by CLECs because, in both cases, scale effects (the necessary scale of entry for viable competition) and the extent of entry itself are affected by changes in such prices. Thus, while it is certainly true that any increase in unbundled switching rates makes self-supplied switching relatively cheaper (thus encouraging its use), it is also clear that increases in the prices of inputs generally do not encourage entry of any sort.

These considerations suggest the following econometric template. As noted, $S^*$ depends on loop rates, unbundled switching rates, and the "minimum necessary market size." Given observations on actual switch deployment by CLECs (conceptually, $S^*$), one may specify a reduced form relationship in which $S^*$ is given as **a** function of loop prices, unbundled switching prices, and other variables relevant to the determination of this minimal market size.[11] These latter magnitudes may be parsimoniously assumed to depend on the average retail revenue in the selected market and on the size of the relevant market. Regulatory policies affect the econometric specification somewhat, as is explained in the next section.

## III. Econometric **Model**

This empirical model focuses on the relationship between CLEC deployed local exchange switching equipment and the rates for unbundled local loops and unbundled local switching. The relationship between wholesale prices and switching facilities deployment is particularly interesting since switch deployment is a primary focus of modern telecommunications policy.

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[11] The relationship between entry and market size is set forth clearly in Sutton (1991, chap 3)

Furthermore, local switching is fertile ground for empirical analysis because state-level data on CLEC deployment of local switching equipment is available, and because UNE prices are established on a state-by-state basis, providing sufficient variability in the data for econometric analysis. In addition, the FCC has limited the availability of unbundled local switching to particular customer-types in certain geographic areas of the Top 50 metropolitan statistical areas.[12] Thus, it is possible to assess how regulatory limitations on access to switching influence switch deployment.

From the Local Exchange Routing Guide ("LERG"), we compute the number of CLEC switches deployed (S) between April 2000 and October 2001 in each of the fifty states and the District of Columbia. Explanatory variables include the price of local loops ($P_L$), the price of unbundled local switching ($P_S$), market size as measured by the number of Bell Company access lines in the state (LINES), and average local service revenue per-line in the state *(RETAIL)*.

In addition, the variable *RESTRICT* measures the percent of population in those metropolitan statistical areas in each state where the availability of unbundled local switching is limited. In the FCC's *UNE Remand* Order, the agency reiterated its position that CLEC access to unbundled local switching ("ULS) is necessary for competition, concluding, "that, in general, lack of access to unbundled local switching materially raises entry costs, delays broad-based entry, and limits the scope and quality of the new entrant's service offerings (¶ 253)." Despite this finding, the FCC chose to remove the unbundled switching obligations of the ILECs for customers with more than three switched access lines in the densest portions (Density Zone 1) of the fifty largest Metropolitan Statistical Areas ("MSA").[13] The rationale for this exclusion was that entrants could serve in a "timely" manner residential and small business consumers at levels of comparable scale and scope by using or deploying their own switching equipment as access to unbundled local switching would allow. The empirical

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[12] The FCC's unbundled local switching restriction allows the incumbent monopolist to either forbid the purchase or raise the price of unbundled switching in most dense portions of the Top 50 MSAs. The restriction did not apply in New York or Texas where state regulations and/or laws prohibited such a restriction. *In re Implementation of the Local Competition Provisions of the Telecommunications Act of 1996, Third Report and Order and Fourth Further Notice of Proposed Rulemaking*, FCC No. 99-238, 15 FCC Rcd 3696 (rel. Nov. 5,1999) ("UNE Remand Order") at ¶ 253-299.

[13] As an additional requirement, the ILEC had to provide access to enhanced extended links ("EELs") in these areas EELs are combinations of loops and transport that "extend" the local loop from one central office to another where the CLEC has collocated equipment. EELs, in theory, reduce the need fur the CLEC to place equipment in every ILEC central office.

model evaluates whether or not the restriction in fact has increased the deployment of switching in these markets by CLECs.

## 1. DATA

As previously mentioned, CLEC switch deployment data is provided by the LERG (January 1999, April 2000, and October, 2001). Bell Company access lines by state are provided by ARMIS From 43-04 (2000 data).[14] Retail price is measured as average revenue per line, and this data **is** provided by the FCC's universal service reports (Federal Communications Commission, Table 5). The percent of population for each state in a restricted, Top 50 **MSA** is computed using Census data.[15] Implicit in the measurement of RESTRICT is that the percent of population in the MSA is highly correlated with the CLECs assessment of the impact of the restriction on its potential market. The restriction applies only to customers with more than three access lines that are also located in the densest portions of the MSA (Density Zone 1, which is a rate zone defined for regulatory purposes). Data measuring the number of customers that fit these criteria are unavailable, *so* this proxy for the scope of the restriction is employed. The econometric model is estimated both with and without the *RESTRICT* variable.

Wholesale prices for loops and unbundled switching are based on state tariffs and interconnection agreements between the ILEC and CLECs. The computation of element costs from this information is both a complex and enormous undertaking. This undertaking was avoided, fortunately, by acquiring summary data on network access prices from a CLEC serving the vast majority of the U.S. market.[16] Loop and switching cost data was provided lor 39 states. Because the other explanatory variables are available for all states, these **39** states make up the final sample.

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[14] The ARMIS data is available online at www.fcc.gov/ccb/armis.

[15] The list of restricted MSAs are provided in Appendix **A** of the *UNE Remand Order*. For MSAs that cross state lines, the population **is** allocated in proportion to the largest cities within **the** MSA. Because the FCC's switching restriction did not apply in New York and Texas, RESTRICT was set equal to zero for these states, A number of states are presently evaluating whether or not to eliminate the restriction (e.g., Georgia and Maryland).

[16] The data was provided **by** Z-Tel Communications, in Tampa, Florida. Z-Tel provides local exchange service using the UNE-Platform (local loops plus local switching/transport) in 39 states (during the time period contemporaneous with the data). Switching costs include local switching and transport, as well as switch related wholesale prices for services such as the daily usage file (which is the file containing usage statistics required for billing for each customer).

## 2. MODEL SPECIFICATION

The econometric equation describing switch deployment is

$$S = \beta_0 + \beta_1 P_L + \beta_2 P_S + \beta_3 LINES + \beta_4 RETAIL + \varepsilon \tag{14a}$$

$$S = \beta_0 + \beta_1 P_l + \beta_2 P_S + \beta_3 LINES + \beta_4 RETAIL + \beta_5 RESTRICT + \varepsilon \tag{14b}$$

where the $\beta$s are the estimated coefficients and $\varepsilon$ is the econometric disturbance term. Because *RESTRICT* is measures somewhat indirectly, Equation (14) has two specifications, one without and one with the *RESTRICT* variable.

The dependent variable (S) is count data (i.e., the data has only discrete values), so we employ the Negative Binomial Regression, which a commonly used alternative to linear least squares regression for count data.[17] Unlike the Poisson regression, which is another popular regression technique for count data, the negative binomial regression does not require that the conditional mean of the data equal the conditional variance. If this assumption is incorrect (i.e., there is *overdispersion* in the data), then the Poisson estimates are invalid and the estimated standard errors are too low (leading to an overstatement of statistical significance) (Gourieroux et al.,). The estimates of the Negative Binomial Regression, however, are not. Further, if overdispersion is not present, then the estimates of the Negative Binomial Regression are identical to those of the Poisson regression.

As a product of the Negative Binomial Regression, and "overdispersion" parameter, $a$, is estimated. The value and statistical significance of this estimated parameter indicates whether or not the Negative Binomial regression is preferred to the Poisson regression, because a noli-zero value of the overdispersion parameter indicates the restrictive assumptions of the Poisson regression are inappropriate (Cameron and Trivedi: 77). If the estimated overdispersion parameter is zero (statistically insignificant), then the Negative Binomial regression is identical to the Poisson regression. Estimates of Models (1) and (2) indicate that overdispersion is present in the data, so the Negative Binomial Regression is the preferred estimation technique for Equation (14).

[17] Ford technical discussion of Negative Binomial and Poisson regressions, *see* Cameron and Trivedi (1998, chap 3)

## 3. RESULTS

'llie results of the Negative Binomial Regression are provided in Table 1, along with descriptive statistics. The dependent variable is measured as the number of CLEC switches deployed in each state between April 2000 and October 2001 (S), during which time the restriction on access to unbundled switching applied.[18] Models (1) and (2), which are based on Equations (14a) and (14b), are estimated as Negative Binomial regressions. An additional model (Model 3) is estimated by least squares, where the dependent variable the natural logarithm of *S*.[19] This alternate specification is a close approximation to the Negative Binomial Regression (as this is made clear by a comparison of the estimated coefficients). While Negative Binomial Regression is theoretically preferred to least squares given the count nature of the dependent variable, the estimates across the specifications are nearly identical.

The likelihood ratio index, a measure of goodness-of-fit for non-linear regressions (such as the Negative Binomial) is 0.76 for Models (1) and (2).[20] The overdispersion parameter, $\alpha$, is statistically significant for both models, indicating that the Negative Binomial Regression is preferred to the Poisson regression. For Model (3), the $R^2$ is 0.59, indicating the model provides for a reasonably good fit, particularly for cross-sectional data. Cameron and Trivedi (1998, p. 89-90) suggest that least squares regressions can be a useful guide for model specification of more complicated count models. Along those lines, two specification tests of Model (3) are conducted. First, the null hypothesis of Ramsey's RESET Test - no specification error - cannot be rejected at anywhere near standard significance levels ($F = 0.33$), which is encouraging. The RESET test is a rather general test of mis-specification, and is capable of detecting omitted variables, endogenous explanatory variables, errors in measurement, and an incorrect functional form for linear models (though it is most sensitive to functional form and omitted variables). Second, the **null** hypothesis of the White Test for heteroscedasticity -- homoscedastic disturbances - was not rejected at

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18 The restriction continues to apply in many states. However, in a number of states - including New York, Illinois, and Texas - the restriction has been eliminated either by legislation or regulatory commission ruling.

19 The least squares regression also was estimated with the RESTRICT variable included. The results were similar to Model (2), and the *RESTRICT* variable was negative and significant. Results are available upon request.

20 For a discussion of goodness-of-fit measures for non-linear regressions and their limitations, see Cameron and Trivedi (1998; 151-8).

standard significance levels ($\chi^2 = 1.05$). Overall, the general specification of the model appears to be sound. at least according to statistical criteria.

For the three models, all explanatory variables are statistically significant **at** the 5% level or better (except for the constant in Model **3).** and the estimated coefficients are robust across specifications. **As** expected, larger markets have more CLEC switches; the coefficient on *LINES* is positive and highly statistically significant (at the 1% level or better). Note that the relationship between access lines and CLEC switches is less than proportional (the estimated coefficient is less than 1.00), indicating that a 10% increase in lines results in about a 4 to 5% increase in switch deployment.[21] Higher revenue per access line also leads to more switch deployment (*RETAIL* is statistically significant and positive), and the elasticity is sizeable (about +2.00). The positive sign on *RETAIL* was expected because higher expected revenues increase the expected profit of entry (ceteris paribus).[22]

Of particular interest are the effects of UNE rates ($P_L$,$P_S$) and the unbundled switching restriction (*RESTRICT*) on CLEC switch deployment. No a priori expectation regarding the effect of the price for unbundled loops or switching on switch deployment was made, given that the theoretical model allows for both positive and negative values (and perhaps a zero value). The regression results indicate, however, that higher loop rates decrease switch deployment; a negative and statistically significant sign on $P_L$ is estimated (with t-statistics larger than 2.00 across all models). The empirical model, by the negative sign on $P_L$, indicates that the entry effect dominates the scale effect. We cannot reject that the estimated coefficient on $P_l$ is equal to -1.00 (via the Wald Test). Thus, assuming a unitary elasticity between switch deployment and loop price is reasonable (i.e., a 10% increase in the loop rate decreases CLEC switch deployment by about 10%).[23]

The theoretical ambiguity between the price for unbundled switching and switch deployment is resolved by the empirical model. The estimated coefficient on the price of local switching ($P_S$) is negative and statistically significant at

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[21] A consistent result is found in Z-Tel Policy Paper No. 4 (2002b).

[22] Theory indicates that existing retail prices may not be a reliable estimate of post-entry prices, so entrants may ignore such prices.

[23] The own-price elasticity of demand for unbundled loops and for unbundled switching are estimated to he -1.7 and -1.00, respectively (Beard and Ford). For loop switching combinations, Ekelund and Ford (2002) and Cline (2002) estimate own-price elasticities of demand in the -2.0 to -3.0 range.

better than the 5% level in each specification. The estimated elasticity is about -0.50, so a 10% increase in the ULS rate decreases CLEC switch deployment by 5%. With respect to the theoretical model, the negative coefficient indicates that, on average, the substitution of unbundled switching for switch deployment is not prevalent at current UNE rates. The entry effect dominates both the scale and substitution effects, so that higher switching rates reduce CLEC switch deployment (ceteris paribus and on average). The result is robust across specifications.

Finally, the sign on *RESTRICT* is negative and statistically significant (at the 5% level) in Model (2), suggesting that the FCC's restriction on access to unbundled switching has impeded rather than encouraged switch deployment. At the sample means, the switching restriction is estimated to have reduced CLEC switching deployment by about 25%. These regression results suggest that the switching restriction impeded rather than encouraged switch deployment.

Given the specification of *RESTRICT*, there is the potential that the variable captures variations in switch deployment across states based factors other than the switching restriction. However, alternative regressions using earlier data indicate that *RESTRICT* has no effect on switch deployment between January 1999 and April 2000, a period prior to the implementation of the restriction. Because the percent of population in a restricted, Top 50 MSA has no effect prior to the implementation of the restriction, but a negative and statistically significant effect after the restriction, it is reasonable to conclude that this measure of the restriction properly captures its effect.[24] In the alternative regression, only market size *(LINES)* and the constant term are statistically significant.

It is also possible that the negative sign on the *RESTRICT* variable captures the possibility that sufficient CLEC switching capacity was deployed in the Top 50 MSAs prior to the restriction, and subsequent to that rule CLECs began deploying switching in smaller markets. To test this hypothesis, the number of switches deployed per access line in each state as of April 2000 was included as a regressor. This additional variable had a positive sign but was not statistically significant; the coefficient on the *RESTRICT* was not much altered and remained statistically significant.[25] The coefficients and t-statistics of the other variables were not materially affected. Given the continued statistical significance of the *RESTRICT* variable in this alternate specification, the alternative hypothesis that

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[24] Results available upon request.

[25] Results available upon request.

CLECs had already deployed enough switching in the Top 50 MSAs prior to the restriction is not supported by the data.

In sum, the econometric results summarized in Table 1 are consistent with economic intuition in that more entry is observed in larger markets and markets with higher prices. The ambiguity with respect to UNE prices is resolved by the negative coefficients on UNE prices (Pi, $P_S$) and availability (*RESTRICT*). In the particular case of switching equipment, this econometric analysis indicates that efforts to encourage the deployment of switching facilities by limiting the availability or increasing the prices of switching UNEs has been (and possibly will be) counterproductive.

## IV. Conclusion

Profit maximizing firms participating in a market economy make "make-or-buy" decisions everyday. While these decisions are of interest to economists in determining what may be an efficient organization of the firm, the "make-or-buy" decision is evaluated differently when the ability to "buy" is mandated and governed by regulation rather than the market, and the ability to "make" is limited substantially by various entry barriers. Such scenarios are increasingly commonplace for the regulated utilities including electricity, gas, and telecommunications, where concepts such as "essential facilities" and "unbundled network elements" are frequently used tools of competition policy.

One common concern in such scenarios is when the ability to "buy" substantially offsets the incentive to "make." In this paper, we evaluated both theoretically and empirically the relationship between "make" and "buy." In our particular construct, where self-supplied and purchased inputs may serve as complements, three sometimes conflicting effects are relevant to the "make-or-buy" decision, of which the substitution effect is only one. Our empirical example considers the deployment of switching facilities by entrants to the local exchange telecommunications markets, and these empirics indicate that the substitution effect is not dominant in this particular case. Of course, the empirical example chosen for our analysis is not necessarily indicative of any other particular case. However, our findings do support the general notion that the substitution effect is not the only relevant consideration, either theoretical or empirical, for policy makers in selecting what inputs to make available to entrants when promoting competition in the utility industries or any industry where mandated access is contemplated or effectuated.

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**Table 1. Negative Binomial Regression Results and Descriptive Statistics (N = 39)**

| | Model (1) Negative Binomial | Model (2) Negative Binomial | Model (3) Least Squares | |
|---|---|---|---|---|
| | Dependent. Variable = $S$ | Dependent. Variable = $S$ | Dependent. Variable = $\ln(S)$ | |
| Variable | Coefficient (t-Stat) | Coefficient (t-Stat) | Coefficient (t-Stat) | Mean (St. Dev) |
| Constant | -6.648 (-2.13)[a] | -10.169 (-3.60)[a] | -5.441 (-1.58) | ... |
| $\ln(P_L)$ | -0.923 (-2.42)[a] | -0.953 (-2.64)[a] | -1.066. (-2.51)[a] | 15.69 (4.60) |
| $\ln(P_S)$ | -0.528 (-2.25)[a] | -0.487 (-2.18)[a] | -0.561 (-2.25)' | 15.53 7.59 |
| $\ln(LINES)$ | 0.392 (3.09)[a] | 0.490 (3.68)[a] | 0.348 (2.65)' | 3,744,347 (4,157,467) |
| $\ln(ARPL)$ | 2.396 (3.30)[a] | 1.917 (2 59)' | 2.31 (2.80)[a] | 33.95 (4.70) |
| $RESTRICT$ | | -0.798 (-1.96)[a] | | 0.30 (0.28) |
| $\alpha$ | 0.295 (5.11)' | 0.268 (5.43)[a] | ... | ... |
| Pseudo $R^2$ | 0.76[c] | 0.76[c] | ... | |
| $R^2$ | ... | ... | 0.59 | |
| RESET F | ... | ... | 0.73 | |
| $S$ | | | | 46.72 (41.59) |

[a] Statistically Significant at the 5% level or better.
[c] Pseudo-$R^2$ is computed using the likelihood ratio index.

# The Demand for Unbundled Elements in Telephony Revisited

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In the December edition of the Atlantic Economic Journal, Drs. Ekelund and Ford [2002] published a paper estimating the own-price elasticity of demand for a combination *of* unbundled elements commonly referred *to* as UNE-P (or UNE-Platform). Using a constant elasticity formulation of the demand curve, Ekelund and Ford ("E-F) estimate an own-price elasticity of –2.7 for UNE-P, indicating that regulated prices for unbundled elements have a potent effect on retail competition in local telecommunications markets.

The sizeable price elasticity has considerable policy implications, suggesting a large entry response to lower wholesale prices. When econometric evidence may influence policy decisions that have sizeable economic consequences, it is important to validate such evidence, both for practical reasons and to assess the impact of particular functional forms on the results. In this note, I conduct an empirical check on the estimates of E-F by evaluating changes in the amount of retail competition using UNE-P arising from recent wholesale price reductions.

The familiar own price elasticity of demand equation $\eta = \%\Delta Q/\%\Delta P$, can be rewritten econometrically as

$$\%\Delta Q = \eta \cdot \%\Delta P + \varepsilon,$$

where $\varepsilon$ is the disturbance term. Using data covering (at least) two time periods, the elasticity term, $\eta$, can be estimated by regressing percent quantity changes on percent price changes, so long as the "distance" between the two periods is short enough that other factors can be treated as constant. While simple, this approach is perhaps consistent with the manner by which policymakers evaluate the effect of price reductions on competitive entry.

This least squares regression (no constant term) is estimated with quantity data for the twelve-month period June 2001 through June 2002 ($\%\Delta Q_{12}$; 35 observations) and the six-month period December 2001 through June 2002 ($\%\Delta Q_6$; 37 observations). Price data is provided by Kovacs *et al.*, [2001, 2002], quantity data by Verizon [2002] for June 2002 data, and Drs. Ekelund and Ford for 2001 data. The estimated equations (standard errors in parenthesis) are

$$\%\Delta Q_{12} = \underset{(0.67)}{-1.83} \cdot \%\Delta P + \varepsilon, \text{ and}$$

$$\%\Delta Q_{6} = \underset{(0.45)}{-1.18} \cdot \%\Delta P + \varepsilon,$$

where the six-month elasticity is –1.18 and a twelve-month elasticity is -1.83. The larger "long-run" elasticity is consistent with theory. Over twelve-months, the own-price elasticity is close to –2.00, and the null hypothesis that the elasticity equals –2.7 cannot be rejected (Wald $\chi^2$ = 1.68). A six-month elasticity of–2.7 is rejected at the **5%** level, but not at the 10% level (Wald $\chi^2$ = 3.26).

This simple test of the reasonableness of the elasticity of demand estimates of E-F affirms their findings, though the full effect of a price reduction on entry may take longer than six months to realize. State regulators that seek to expand competitive choice in retail local telecommunications markets by reducing wholesale prices can expect to see elastic responses of service provided over the combination of unbundled elements called UNE-P.

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